Assessing a Perfect European Optimum Currency Area: a Common Cycles Approach

Michel Beine [1], Bertrand Candelon [2] & Alain Hecq [3] [4]

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[1] Ministry of the Walloon Region and University of Lille II, CADRE, 1, place Déliot, 59324 Lille, France; Beine@hp-sc.univ-lille2.fr.

[2] Institut für Statistik und Ökonometrie, Humboldt University zu Berlin, Spandauer Strasse 1, 10178 Berlin, Germany; candelon@wiwi.hu-berlin.de

[3] Maastricht University, Department of Quantitative Economics, P.O.Box 616, 6200 MD Maastricht, The Netherlands; a.hecq@ke.unimaas.nl.

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Abstract

In this paper we introduce a new definition for an optimum currency area (OCA) which is more restrictive than the previous ones. Indeed, using both a cointegration and a common cyclical feature analysis in a VAR(p) framework, a set of countries is said to constitute a perfect OCA if the short-run dynamics is perfectly correlated while long-run relationships are not constrained. Using seasonally unadjusted industrial production indices for the period 75:M1 to 97:M4, we show that European countries are not sufficiently related to fit our definition.

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1 Introduction

Like it or not, the theory of optimum currency areas (OCA) remains the implicit reference framework to assess the real consequences of monetary integration. Initiated by Mundell (1961), the OCA approach has developed a set of criteria aimed at investigating whether a particular geographic zone will gain from adopting a single currency. These criteria may depend on labor mobility and trade openness (Mc Kinnon, 1963), industrial diversification (Kenen, 1969), wage and price flexibility (Blanchard and Muet, 1993) as well as fiscal federalism (Sachs and Sala-I-Martin, 1992). The project of a European Monetary Union (EMU) has led to a renewal of this old Keynesian approach. On the theoretical side, general equilibrium analysis (Bayoumi, 1994; Ricci 1997) has fulfilled the lack of modelling which characterizes the traditional core of the theory[1].

On the empirical side, the OCA approach has been applied to investigate whether Europe may be considered as an optimum currency area. There is nowadays a widespread consensus between economists to state that the answer to this question is negative and that only a subset of countries will gain from giving up their national currencies. However, the precise size and configuration of this restricted monetary union remains a matter of debate. Indeed, direct applications of the above mentioned criteria turn out to be cumbersome because different empirical counterparts of the theoretical variables may yield quite different outcomes and consequently different implications for the composition of the future EMU. This is for instance the case for the criteria of trade openness and industrial diversification. As a consequence, the empirical analysis have focused on an implicit but overwhelming criterion : the importance of real asymmetric shocks affecting the members of the monetary union. As shocks are not observable, it is hardly surprising that econometric methods have been extensively used in order to measure their asymmetry degree.

It comes out that this important literature faces some problems. One important problem concerns the use of bivariate statistical methods to determine the OCA which is, by definition, a multi-country concept. Another crucial point concerns the distinction between the short-run and the long-run dynamics. In this paper, we try to fulfil those lacks by analyzing in a multivariate VAR(p) framework both long-run and short-run comovements between economic indicators of a set of European countries. Economic theory often deals in this matter with

[1]See Mélitz (1995) on this point.

long-run comovements via convergence analyses between the levels of GDP or other economic indicators. Those types of comovements have been largely studied in empirical works[2], through cointegration analysis. The analysis of the short-run fluctuations has been a matter of fewer investigations. However, in the light of the OCA approach, it is more relevant to study whether a monetary union is sustainable on the basis of short-run relationships .

In this study, we check the existence of a core of countries within the EC. We introduce a new empirical definition for an OCA which, in some way, is more restrictive than the previous ones. Using a framework developed in Hecq and al. (1997a,b) for VAR(p) with common cyclical feature and cointegration, a set of countries is said to constitute a perfect OCA if the short-run dynamics is perfectly correlated while the long-run relationships are not constrained. More precisely, we define this core by the set of countries for which exogenous shocks adjust instantaneously, and for which there exists only one common cycle. Indeed, the number of common cycles constitutes a reliable indicator of short-run coordination. If there is less common cycles than countries, there is a strong signal of policy and/or economic coordination.

The paper is organized as follows. Section 2 presents a short survey of the empirical literature of OCA and emphasizes the drawbacks of most of existing studies. Section 3 presents the econometric methodology relying on the recent developments of the common cyclical features literature. The empirical investigation is reported in section 4. Section 5 concludes.

2 Existing Empirical Approaches of OCA

The theory of OCA focuses on the stabilization cost induced by the loss of the nominal exchange rate as a policy instrument. The size of this stabilization cost has been found to depend on a set of structural features of the economies thought to adopt a common currency.

2.1 Direct Applications of OCA

In order to determine the exact composition of the future EMU, some authors have proposed to apply directly some of the structural criteria. It comes out that these direct measures often lead to divergent conclusions about the composition of the future EMU. An important reason lies in the diversity of the empirical measures.

[2] See for instance Bernard and Durlauf (1995).

Countries	Au	Be	Dk	Sp	Fin	Fr	Gr	Irl	I	Nl	Pr	UK	Sw
Gros(1996)	5	2	4	12	8	10	13	1	11	3	7	9	6
Pisani-Ferry(1997)	2	1	10	3	12	9	6	13	4	7	5	8	12

Table 1: The European OCA on the Basis of Trade Openness

Table 1 provides an illustration for one of the most important criteria, i.e. trade openness. Two different measures of openness are reported, in terms of ranks. The first one, proposed by Gros (1996), is the share of exports towards the European Union[3]. The second one reported by Pisani-Ferry (1997) relies on the importance of trade flows with the so-called EMS Core[4]. Application to Europe 12 leads to a ranking of European candidates and in turn to the desirable composition of the future EMU. From Table 1, it comes out that quite opposing results are obtained for some countries. For instance, the first measure suggests that on a relative basis, trade openness does not favor Spain or Italy's inclusion while it does on the basis of the second indicator[5]. Though less clear-cut, the same contradiction applies for the criterion of industrial diversification.

2.2 Shocks Asymmetry

In order to cope with the problem of diverging measures, most empirical OCA studies have relied on an implicit but central criterion in the core of the theory, i.e. the importance of asymmetric (real) shocks. In the absence of shocks of this nature, no adjustment tool is required and the cost of losing the nominal exchange rate is expected to be outweighed by the benefits induced by a common currency[6]. The importance of this criterion stems from the fact that the alternative stabilization channels at the European level are found in a large extent to be ineffective. One may identify four main channels. The first one, emphasized by Mundell (1961), is labor mobility. All studies (see for instance OECD 1999) conclude that transnational labor mobility is very limited[7]. Furthermore, the American experience in the field of migration suggests that the response to

[3]Exports towards European Union 15 as a percentage of GDP; the ranking is adjusted in order to account for the inclusion of Germany.

[4]Trade flows with EMS Core minus flows with the rest of the world as a percentage of GDP.

[5]Of course, this does not mean that Spain or Italy are excluded or included in the OCA. One of the general weakness of the OCA criteria taken on an induvidual basis is that they do not involve thresolds. This is one of the reasons why the empirical literature has developed OCA indices (Bayoumi and Eichengreen, 1997).

[6]Traditionally, these benefits refer to microeconomic efficiency gains. De Grauwe (1996) for instance offers an evaluation of the expected respective importance of these benefits.

[7]Furthemore, Eichengreen (1993) shows that even within a country, labour mobility remains also limited.

income discrepancies can happen after an important lag (30 to 40 years in the case of the South-North black migrations in the United states). As a result, labor mobility cannot be considered as a stabilization tool. The second traditional stabilization channel is (real) wage flexibility. For stabilization analysis, wage flexibility encompasses two dimensions. The first one is wage differentiation across sectors and regions. The second one is the wage dependence on economic fluctuations. In both dimensions, this flexibility has been found to be very low (Bean 1994, Vinals and Jimeno, Abraham 1996). Furthermore, even though increased wage differentiation between countries, regions or sectors appears desirable, Blanchard and Katz (1992) show that its stabilization effect is quite limited. The third stabilization mean goes through national discretionary fiscal policies. Their use is nevertheless highly restricted either from the need to reduce the high debt ratios inherited from the 80's (like in Belgium or Italy for instance), either by the provisions of the Stability and Growth Pact which lead to pro-cyclical policies. Finally, while budgetary transfers in favor of poor regions have been provided for redistribution purposes, the scope for a fiscal federalism system aiming at stabilizing temporary depressed countries is very limited in Europe. To sum up, since almost no stabilization instrument is left for EMU members, the extent of asymmetric shocks (or generally speaking the degree of cycle disynchronisation) is the key measure of the cost of EMU.

As shocks are by definition not observable, it is hardly surprising that these analyses have extensively relied on econometric methods. Table 2 presents a sample of this extensive literature. In particular, the econometric method and the implication in terms of which countries expected to form the EMU (the so-called " core ") are reported. Finally, whether the analysis relies on a bivariate approach or on a multivariate approach is a reported feature. Here, bivariate should be understood as country-to-country correlations without accounting for the relationship with other countries. By contrast, multivariate refers to a multi-country approach in which all the relationships within a specific group ("the OCA" for instance) are accounted for.

Some comments are in order. (i) First, there is no widespread agreement about the precise composition of the European EMU. Except for Germany and its small neighbors (Belgium and the Netherlands) which are considered as natural participants, the inclusion of the other European countries remains a matter of debate. The general conclusion emerging from this literature is that there is a distinction between a core of countries facing similar disturbances and a periphery in which idiosyncratic shocks are found to be important. This general assessment has nevertheless been challenged by Rubin and Thygesen (1996) who conclude in favor of a

Paper	Method	Approach	Core
von Hagen and Neumann(1994)	Real exchange rate variability	bivariate	Ge, Au, Be, Nl,
DeSerres and Lalonde (1995)	Real exchange rate variability	bivariate	Ge, Be, Nl
Helg and al. (1995)	Sectorial Analysis/VAR	bivariate	Ge, Be, Nl, Dk, Fr, UK
Bayoumi and Eichengreen (1993)	Structural VAR	bivariate	Ge, Be, Nl, Dk, Fr
Artis and Zhang (1995)	Cyclical Components Identification	bivariate	Ge, Be, Fr, Nl, Sp, Pr, It
Beine and Hecq (1997)	Codependence in VMA	bivariate	Ge, Be, Nl
Rubin and Thygesen (1996)	Codependence in VAR	multivariate	Ge, Fr, Be, Nl, Dk, Fi

Table 2: Optimal Composition of EMU and Shocks Asymmetry

widespread sustainable monetary union with nine countries (including countries like Italy, the UK or Sweden, traditionally viewed as outsiders). (ii) Secondly, those analysis make use of quite different statistical methods, sometimes belonging to opposite econometric paradigms. For instance, some approaches consider trends of the exclusively deterministic type (Artis and Zhang, 1995) while others view these trends as stochastic (Rubin and Thygesen, 1996). (iii) Finally, one may identify some shortcomings characterizing this important empirical literature. These are sketched out here below.

2.3 Weaknesses in the Existing Literature

The first problem concerning this empirical literature is that most of the analyses rely on bivariate methods[8]. For instance, in the Bayoumi and Eichengreen (1993) analysis which may be considered as an important reference study, the degree of asymmetry is measured through the correlations of identified supply and demand shocks (on the basis of the Blanchard and Quah (1989) decomposition) with those of a reference country, i.e. Germany in this case[9]. Such an approach however neglects an important point which has been emphasized by some authors (see for instance De Grauwe, 1996). An OCA is by definition a multi-country concept and the degree of asymmetry of each country is conditional on the precise composition of the currency union.

[8] A noticeable exception is Rubin and Thygesen (1996).

[9] In an extension of this paper, Bayoumi and Eichengreen (1996) report all cross-country correlations and infer the European currency area. These measures remain however based on bivariate correlations.

The entrance into the union of a new candidate may therefore modify the importance of asymmetric shocks of each country, this change depending on the trade linkages with this country. In turn, this implies that some EMU countries may support the inclusion of some trade partners while disregarding other "candidates". For instance, if Ireland was to belong to EMU, one may expect it to support the inclusion of the UK because these economies share similar business cycles. Accounting for this important issue requires the use of multivariate statistical methods.

Another problem is the distinction between the short and the long-run dynamics. In the context of OCA, this distinction is crucial. Indeed, OCA theory emphasizes the role of the nominal exchange rate as a stabilization instrument. In the long-run, nominal exchange rate variations are ineffective in correcting differential developments in costs or demand [10]. The Belgian experience of 1982 indicates that realignments of the nominal exchange rate may be successful during 1 or 2 years to dampen an important recession. After this period, structural policies on the labor market are needed to enhance the stabilization gains obtained from the change in the nominal exchange rate. Likewise, the EMU implies the loss of discrete national monetary policies. Recent estimations from Barran, Coudert and Mojon (1997) indicate that the impact on GDP of a change in the short-run interest rates lasts for about one year and a half[11]. Unfortunately, the distinction between short and long-run shocks is seldom explicitly accounted for. For instance, Bayoumi and Eichengreen (1993) estimate bivariate VARs in first differences and consequently make no distinction between stochastic trends and common cycles. Our estimation procedure allows to identify the short-run dynamics from the long-run one.[12]

A final important issue concerns the way one assesses, in a coherent fashion, the degree of asymmetry in the econometric model. Along the lines mentioned here above, this measure should reflect the influence, for each country, of short-run, symmetric shocks jointly determined by the dynamics of all countries considered in the OCA. The importance of the common cycle provides a relevant measure of this degree of asymmetry. This feature stands in opposition with some other measures and particularly with the one proposed by Rubin and Thygesen (1996). Moreover, in order to assess the robustness of the results, it is important to consider models in which no a priori restriction is imposed, as it is done in some papers like Bayoumi and

[10] See De Grauwe (1994, chap. 2) on this particular point.

[11] Transmission mechanisms appear quite different with respect to the timing of the full effect .This ranges from less than 6 months in the UK to 30 months in Germany. The average lies at around one year.

[12] Furthermore, given the choice of the lags in our VAR model, the definition of the short run process (10 months) is consistent with the reported duration of stabilisation policies (through exchange rate management or monetary policy).

Eichengreen (1993) or, to a less extent, Rubin and Thygesen (1996)[13].

In this paper, we adopt an econometric specification aimed at simultaneously dealing with these issues. We therefore use a common cycle-common trend approach relying on the recent developments of the common features literature (Engle and Kozicki, 1993; Vahid and Engle, 1993). In particular, we consider a multivariate VAR model with cointegration and common cyclical features which explicitly distinguishes the short-run dynamics (driven by the common cycle(s)) and the long-run evolution (driven by the common trend(s)). We then introduce an explicit definition of a "Perfect OCA":

Definition 1 *A set of countries constitute a "Perfect OCA" iff there exists only one (weak form) synchronous common cycle and their common features vectors span a space excluding the presence of asymmetric shocks.*

We develop these statistical concepts in the next section.

3 The Methodology

3.1 VAR Models with Cointegration and Common Features

Let us consider a Gaussian Vector Autoregression of finite order p (VAR(p)) model for an n-dimensional I(1) vector time series $\{x_t, t = 1, \ldots, T\}$:

$$x_t = m + \sum_{i=1}^{p} \Pi_i x_{t-i} + \varepsilon_t, \qquad t = 1, \ldots, T \tag{1}$$

with fixed initial values of $x_{-p+1}, ..., x_0,$ m is a vector of constants and where ε_t is a n-dimensional homoskedastic Gaussian mean innovation process relative to $(x_{t-1}, x_{t-2}, \ldots, x_1)$ with nonsingular covariance matrix Ω. Let L denote the lag operator and define $\Pi(L) = I_n - \sum_{i=1}^{p} \Pi_i L^i$. We further assume that the process x_t is cointegrated of order (1,1) (see Johansen, 1995). The $rank(\Pi(1)) = r$, $0 < r < n$, so that $\Pi(1)$ can be expressed as $\Pi(1) = -\alpha\beta'$, with α and β both $(n \times r)$ matrices of full column rank r. The columns of β span the space of cointegrating vectors, and the elements of α are the corresponding adjustment coefficients or factor loadings.

[13] In fact, due to the methodology adopted (Vahid and Engle 1993), Rubin and Thygesen (1996) implicitly impose an upper bound to the number of common cyclical features vectors and thus a lower bound to the number of common cycles. This will be clarified in Section 3 in which the difference between the strong and the weak form reduced rank structures is exposed.

Decomposing the matrix lag polynomial $\Pi(L) = \Pi(1)L + \Gamma(L)(1-L)$, and defining $\Delta = (1-L)$, we obtain the vector error correction model:

$$\Delta x_t = m + \alpha\beta' x_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta x_{t-i} + \varepsilon_t, \qquad t = 1, \dots, N, \tag{2}$$

where $\Gamma_0 = I_n$, $\Gamma_j = -\sum_{k=j+1}^{p} \Pi_k$ $(j = 1, \dots, p-1)$ and $\alpha\beta' = -\Pi(1) = -(I_n - \sum_{j=1}^{p} \Pi_j)$.

Serial correlation common feature (SCCF, see Engle and Kozicki, 1993) holds for the VECM (2), if there exists a $(n \times s)$ matrix $\tilde{\beta}$, whose columns span the cofeature space, such that $\tilde{\beta}'(\Delta x_t - m) = \tilde{\beta}'\varepsilon_t$ is a s-dimensional vector mean innovation process with respect to the information available at time t. Consequently, serial correlation common features hold if the cofeature matrix $\tilde{\beta}'$ satisfies the following two conditions:

$$\textbf{Assumption 1:} \qquad \tilde{\beta}'\Gamma_j = 0_{(s\times n)}, \qquad j = 1 \dots p-1 \tag{3}$$

$$\textbf{Assumption 2:} \qquad \tilde{\beta}'\Pi(1) = -\tilde{\beta}'\alpha\beta' = 0_{(s\times n)} \tag{4}$$

Assumption 1 implies that $\tilde{\beta}'$ must lie in the intersection of the left null spaces of the matrices describing the short-run dynamics $\Gamma_j = 0$, $j = 1 \dots p-1$. Given that $\Gamma_j = -\sum_{k=j+1}^{p} \Pi_k$, $j = 1, \dots, p-1$ and $-\Pi(1) = -(I_n - \sum_{j=1}^{p} \Gamma_j)$, Assumption 2 implies that $\tilde{\beta}'(I_n - \Pi_1) = 0_{(s\times n)}$, e.g. Π_1 must have eigenvalues equal to one with multiplicity equal to s and the corresponding eigenvectors must lie in the intersection of the left null spaces of the Γ_j matrices. Cointegrated VAR models satisfying both Assumptions 1 and 2 are considered in detail in Vahid and Engle (1993). Considering models that satisfy either both Assumptions 1 and 2 or Assumption 1 only, Hecq, Palm and Urbain (1997a,b) introduce the two following definitions:

Definition 2 (Strong Form Reduced Rank Structure): *If in addition to cointegration both Assumptions 1 and 2 hold, the implied reduced rank structure of the VECM (2), will be labelled a strong form reduced rank structure (SF). Under SF, there exists a $(n \times s)$ matrix $\tilde{\beta}$, whose columns span the cofeature space, such that $\tilde{\beta}'(\Delta x_t - m) = \tilde{\beta}'\varepsilon_t$ is a s-dimensional vector mean innovation process with respect to the information available at time t.*

Definition 3 (Weak Form Reduced Rank Structure): *If in addition to cointegration only Assumption 1 holds, the implied reduced rank structure of the VECM (2), will be labelled a weak form reduced rank structure (WF). Under WF, there exists a $(n \times s)$ matrix $\tilde{\beta}$, whose columns span the cofeature space, such that $\tilde{\beta}'(\Delta x_t - m - \alpha\beta' x_{t-1}) = \tilde{\beta}'\varepsilon_t$ is a s-dimensional* *vector mean innovation process with respect to the information available at time t.*

It is important to note an important difference between SF (or SCCF) and WF[14] since in the latter case both the possible number and the nature of the common features change: s may be greater than $n-r$ but has to remain $\leq n-1$ and the corresponding $n-s$ common dynamic factors consist of linear combinations of the lagged first differences only. Also note that in the WF, we consider linear combinations of the first differenced I(1) variables that are predictable in the long-run, i.e. at low frequencies. The SCCF rules out predictability at any frequency.

Example 2 *In order to stress the difference between the SF and WF models, let us take a cointegrated VAR of order $p=2$, written in its VECM form, for the GDP of two countries. We get for $t=1\dots T$:*

$$\begin{bmatrix}\Delta y_{1t}\\ \Delta y_{2t}\end{bmatrix}=\begin{bmatrix}\mu_1\\ \mu_2\end{bmatrix}+\begin{bmatrix}\lambda\phi_{21} & \lambda\phi_{22}\\ \phi_{21} & \phi_{22}\end{bmatrix}\begin{bmatrix}\Delta y_{1t-1}\\ \Delta y_{2t-1}\end{bmatrix}+\begin{bmatrix}\alpha\\ 1\end{bmatrix}[\beta_1 \quad \beta_2]\begin{bmatrix}y_{1t-1}\\ y_{2t-1}\end{bmatrix}+\begin{bmatrix}\varepsilon_{1t}\\ \varepsilon_{2t}\end{bmatrix} \tag{5}$$

where μ_1 and μ_2 are constant drift terms, ε_{1t} and ε_{2t} are white noise processes with non singular covariance matrix Ω. (β_2/β_1) is the long-run elasticity if country 1 GDP is the normalizing variable. Because the first row of the autoregressive coefficients matrix is a linear combination of the second one, this matrix is also of reduced rank and the system may be written as :

$$\begin{bmatrix}\Delta y_{1t}\\ \Delta y_{2t}\end{bmatrix}=\begin{bmatrix}\mu_1\\ \mu_2\end{bmatrix}+\begin{bmatrix}\lambda\\ 1\end{bmatrix}[\phi_{21} \quad \phi_{22}]\begin{bmatrix}\Delta y_{1t-1}\\ \Delta y_{2t-1}\end{bmatrix}+\begin{bmatrix}\alpha\\ 1\end{bmatrix}[\beta_1 \quad \beta_2]\begin{bmatrix}y_{1t-1}\\ y_{2t-1}\end{bmatrix}+\begin{bmatrix}\varepsilon_{1t}\\ \varepsilon_{2t}\end{bmatrix} \tag{6}$$

Consequently, in SF model, i.e. the original formulation proposed by Engle and Kozicki (1993), the long-run and short-run matrices share the same left null space. This corresponds to $\alpha=\lambda$ in system (6). In this case, there exists a common feature vector $\tilde{\beta}'=[1 \quad -\lambda]$ such as premultipliying expression (6) by $\tilde{\beta}'$ yields a white noise. In the less restrictive WF model, $\alpha\neq\lambda$ and then a linear combination of first differences in deviation from the long-run equilibrium is white noise.

Common features relationships yield information relative to the short-run comovements. Those relationships may come from economic theory (relative purchasing power parity) or from more empirical facts called "stylized facts" (convergence) and give the common dynamic factor

[14]Notice that a mixed model which presents both WF and SF is also defined in Hecq and al. (1997a).

driving the dynamics of the system, i.e. $\phi_{21}\Delta y_{1,t-1} + \phi_{22}\Delta y_{2,t-1}$ in the WF case for instance[15]. Notice that because the linear combination gives a white noise, the cycle is said to be synchronous while it will be said unsynchronous if a moving average component remains in the process (see Vahid and Engle, 1997). The orthogonal complement of the space $\tilde{\beta}$, labelled $\tilde{\beta}_{\perp}$ is such as $\tilde{\beta}^{'}\tilde{\beta}_{\perp} = 0_{s\times n}$, gives the factor loading of the common factor in the equations. For instance, $\tilde{\beta}_{\perp} = [\lambda \quad 1]^{'}$ in system (6). The implications of these two classes of models are discussed in Hecq, Palm and Urbain (1997a,b) where inferential issues as well as permanent-transitory decompositions are investigated. We summarize some results in the next section.

3.2 Testing Procedure

Within what is a two steps approach[16], there are two ways to treat the estimation problem. The first way is to carry out a canonical correlation analysis between the variables on one hand and the set of instruments on the other hand. The non-significant canonical correlations reveal the existence of linear combinations which yield white noise processes. The second approach relies on instrumental variables or on the generalized method of moments. In this framework, a test of serial correlation common features is a test of overidentifying restrictions. The use of canonical correlation estimation has the advantage that results do not rely on the choice of a dependent variable. Moreover it is more convenient when testing for more than one common feature vector.

All the test statistics we consider here are based on the null that the s smallest canonical correlations are zero. Let us define the$T \times n$ matrices $W_1 = \Delta X = (\Delta x_1, \dots, \Delta x_T)^{'}$, $X_{-1} = (x_0, \dots, x_{T-1})$, $Z_1 = \Delta X^*$, with ΔX^* being the LS residuals from the multivariate regression of ΔX on $X_{-1}\beta$, and the $T \times (n(p-1)+r)$ matrix $W_2 = (Z_2, X_{-1}\beta)$ with Z_2 being the $T \times n(p-1)$ matrix $(\Delta X_{-1}, \dots, \Delta X_{-p+1})$. Under SF, we obtain the sequence of common feature likelihood ratio tests statistics for $H_0 : rank(\tilde{\beta}) \geq s$ against $H_A : rank(\tilde{\beta}) < s$ is $\xi_S = -T\sum_{i=1}^{s} \ln(1-\lambda_i)$, for $s = 1, \dots, n-r$, where $0 \leq \lambda_1 \leq \dots \leq \lambda_{n-r}$ are the ordered eigenvalues of the symmetric matrix $(W_1W_1^{'})^{-1/2}W_1W_2^{'}(W_2W_2^{'})^{-1}W_2W_1^{'}(W_1W_1^{'})^{-1/2}$. For known r and β, under the null the test statistic ξ_S is asymptotically χ^2-distributed with $s(n(p-1)+r) - s(n-s)$ degrees of freedom. Under WF, the sequence of common feature likelihood ratio tests statistics is for

[15]To extract the common cycle, one needs a particular trend/cycle decomposition which uses both cointegrating and common cyclical feature restrictions.

[16]In the first step, one checks for the presence of cointegrating relationships and then fixes it to carry out the common feature analysis in a second step.

$H_0 : rank(\tilde{\beta}) \geq s$ against $H_A : rank(\tilde{\beta}) < s$ is $\xi_W = -T\sum_{i=1}^{s} \ln(1-\tilde{\lambda}_i)$, for $s = 1, \ldots, n-1$, where the ordered eigenvalues of the symmetric matrix $(Z_1 Z_1^{'})^{-1/2} Z_1 Z_2^{'} (Z_2 Z_2^{'})^{-1} Z_2 Z_1^{'} (Z_1 Z_1^{'})^{-1/2}$ are $0 \leq \tilde{\lambda}_1 \leq \ldots \leq \tilde{\lambda}_{n-1}$. For known r and β, under the null the test statistic ξ_S is asymptotically χ^2-distributed with $s(n(p-1)) - s(n-s)$ degrees of freedom.

For a given r and for each s, a likelihood ratio test statistic for the null hypothesis of a SF against the alternative of a WF, for each possible common feature rank $i = 1 \ldots n-r$, is given by $\xi_{SW} = -T\sum_{i=1}^{n-r} \ln\{(1-\lambda_i)/(1-\tilde{\lambda}_i)\}$, where ξ_{SW} has an asymptotic $\chi^2_{(sr)}$ distribution.

4 Empirical Investigation

The data we use as well as their statistical characteristics do have a deep impact on the common cyclical feature analysis. To see that, one should keep in mind two important points. First, Beine and Hecq (1999) illustrated through Monte Carlo experiments that temporal aggregation yields a large decrease in the power of common feature statistics. Ignoring this point may lead to finding more common feature vectors than the ones that could exist in a model representing the behavior of economic time series. Secondly, it is shown in Hecq (1998) that seasonal adjustment induces both huge size and power distortions. As a consequence, nothing could be inferred from seasonally adjusted data for common feature test statistics. Using seasonally adjusted quarterly data for real gross domestic product and unemployment rate, we find for both variables three cointegrating vectors in an "a priori" core composed of Germany, Belgium, The Netherlands, France, Austria and Denmark[17]. Conditioning on those long-run relationships, two common feature vectors are found for the unemployment rate while one or two for real GDP. All of them are meaningless and may just reflect the fact that the time series are inappropriate for our study.

Consequently, we focus in our analysis on seasonally monthly unadjusted industrial production indexes for the period ranging from 1975:M1 to 1997:M4. Monthly industrial production indexes have already been analyzed within a common feature framework by Rubin and Thygesen (1996). However, their analysis relied on seasonally adjusted data even if they carry out regressions with seasonal dummies. In order to save space, we only present the outcomes concerning a "starting Core" of European countries made of Germany, Belgium, The Netherlands, Austria and France. Adding other countries does not modify the conclusions we draw from this analysis. Table 3 gives the results of monthly unit root tests both at the zero and the seasonal frequencies

[17] In order to save space, these results are not reported here. They are nevertheless available upon request.

	Model	lags	π_1	π_2	$\pi_3 \cap \pi_4$	$\pi_5 \cap \pi_6$	$\pi_7 \cap \pi_8$	$\pi_9 \cap \pi_{10}$	$\pi_{11} \cap \pi_{12}$
Be	c, sd	1 to 8	−.61	−1.60	5.52	10.96∗	12.49∗	10.17∗	12.86∗
	c, sd	7, 8	−.62	−2.15	7.84∗	18.99∗	31.03∗	10.16∗	21.58∗
	c, sd, t	1 to 8	−2.23	−1.61	5.71	10.63∗	12.01∗	9.84∗	12.62∗
	c, sd, t	7, 8	−2.18	−2.15	8.44∗	18.88∗	31.19∗	9.46∗	20.53∗
Nl	c, sd	1 to 12	.48	−2.31	6.38∗	9.48∗	2.45	6.67∗	6.34∗
	c, sd	2, 5, 12	.58	−3.41∗	6.54∗	18.07∗	2.91	8.43∗	10.03∗
	c, sd, t	1 to 12	−2.22	−2.34	6.38∗	9.67∗	2.56	6.78∗	6.63∗
	c, sd, t	2, 5, 12	−2.03	−3.42∗	18.54∗	19.36∗	3.40	8.79∗	10.81∗
Au	c, sd	1 to 6	−.38	−1.66	4.95	6.83∗	15.37∗	14.61∗	16.57∗
	c, sd	1, 3, 4, 5, 6	−.38	−1.81	4.79	7.42∗	15.25∗	14.62∗	16.59∗
	c, sd, t	1 to 6	−3.14	−1.70	4.68	6.51∗	12.54∗	15.06∗	16.62∗
	c, sd, t	1, 3, 4, 5, 6	−3.13	−1.87	4.47	7.11∗	12.36∗	14.99∗	16.62∗
Ger	c, sd	1 to 11	−1.02	−3.16∗	6.57∗	11.68∗	3.61	12.86∗	8.17∗
	c, sd	1, 2, 3, 5, 10, 11	−1.04	−3.56∗	9.82∗	13.02∗	3.96	12.77∗	9.57∗
	c, sd	1 to 11	−1.83	−3.13∗	6.48∗	11.45∗	3.56	12.64∗	8.24∗
	c, sd	1, 2, 3, 5, 10, 11	−1.77	−3.52∗	9.61∗	12.84∗	4.09	12.32∗	9.85∗
Fr	c, sd	1 to 12	−1.13	−2.18	2.98	4.04	1.06	9.42∗	2.88
	c, sd	1, 2, 3, 6, 11, 12	−1.15	−2.55	3.53	4.44	1.08	11.72∗	3.09
	c, sd	1 to 12	−1.73	−2.20	3.03	4.03	1.05	9.32∗	2.87
	c, sd	1, 2, 3, 6, 11, 12	−1.77	−2.57	3.57	4.39	1.07	11.36∗	3.14

Table 3: Seasonal Unit Root Tests

(see Franses, 1991, 1996). We report tests with a constant and seasonal dummies (c,sd) as well as those with a constant, seasonal dummies and a deterministic trend (c,sd,t). For these two regressions, Table 3 also gives unit root test statistics, with two ways for whitening the residuals. In the first one, we augment the regression with 1 to k lags, where k is the last significant lag using a $t-$statistic criterion. We also present results when we dropped the non significant lags. A * means that the test statistic rejects at a 5% level the null hypothesis of a unit root at that specific frequency. It emerges from the table that all time series present a unit root at the zero frequency while they have different seasonal unit roots. Consequently, a seasonal cointegration analysis turns out to be difficult. We assume here that there are no seasonal cointegrating vectors and we filter the seasonal unit roots in order to study cointegration at the non seasonal frequency. As $1 - L^{12} = (1 - L)(1 + L)(1 + L^2)(1 + \sqrt{3}L + L^2)(1 - \sqrt{3}L + L^2)(1 + L + L^2)(1 - L + L^2)$, we apply the filter $(1 - \sqrt{3}L + L^2)$ for Germany and The Netherlands, $(1 + L)$ for Belgium, $(1 + L + L^2 + L^3)$ for Austria and $(1 + L^3 + L^6 + L^9)$ for France.

Formally, both Johansen's traditional maximum likelihood statistics would retain only one cointegrating vector for those transformed data. However, both a graphical analysis and a check of the largest eigenvalues of the companion matrix conclude in favor of two vectors. The VAR model contains a linear trend constrained in the long-run, and a constant, seasonal dummies and an outlier in 1984:M6 (for Germany) unconstrained in the short-run. The lag length is chosen equal to 10 according to the Hannan Quinn criterion. We obtain one cointegrating vector if we reduce this core to only three countries, namely Germany, Belgium and The Netherlands. The following matrices report the normalized cointegrating vectors for both cases.

$$\beta' x_t = \begin{bmatrix} 1 & 0 & -.078 & \text{-}.102 & -.036 & .00072 \\ 0 & 1 & -.078 & .113 & -.112 & .00011 \end{bmatrix} \begin{bmatrix} \text{Ger} \\ \text{NL} \\ \text{Au} \\ \text{Fr} \\ \text{Be} \\ \text{Trend} \end{bmatrix}$$

$$\beta' x_t = \begin{bmatrix} 1 & .213 & -.281 & .00028 \end{bmatrix} \begin{bmatrix} \text{Ger} \\ \text{NL} \\ \text{Be} \\ \text{Trend} \end{bmatrix}$$

Fixing those vectors, we obtain test statistics presented in Section 3.2. We test for zero canonical correlations between Δx_t and $\left\{\Delta x'_{t-1}, ..., \Delta x'_{t-9}\right\}'$, where both sets have been concentrated on $\beta' x_{t-1}$, the constant and the dummies. Table 4 gives the WF common feature test statistics for these two sets of countries for the whole period (1975M1-1997M4) as well as for two subperiods. The first one starts in 1983M3, with the second phase of the EMS and the second subperiod ranges from 1986M9 to the end of the sample, corresponding to the so called new EMS[18]. The eigenvalues as well as the number of degrees of freedom of the χ^2 distribution under the null are also presented. One sees that we reject the presence of any weak form common feature vectors and obviously also $n-1$ of them.

For both sets of countries, it is thus impossible to isolate a unique common cycle driving the short-run dynamics of the series. Consequently, using the industrial production indexes to summarize economic activity, we are unable to determine what we called a "Perfect OCA". At least, for this variable, codependence could be used in order to determine whether the adjustment speed is lower for some combination of countries than for the countries individually.

To illustrate this point further, we have considered, following the approach developed by Vahid and Engle (1997), the presence of codependent cycles (or non-synchronous cycles) for the whole period (minor modifications appear for subperiods). Table 5 gives the p-values of the test statistics for different order of codependence (denoted b). *Cod(b)* means there exists a

[18]The cointegrating vectors are of course also computed for these subperiods.

		Ge,Be,Nl,Fr,Au			Ge,Be,Nl			
		$\tilde{\lambda}_i$	$-T\sum\ln(1-\tilde{\lambda}_i)$	df		$\tilde{\lambda}_i$	$-T\sum\ln(1-\tilde{\lambda}_i)$	df
75M1-97M4	$s=1$	.31	91.64	41	$s=1$	.49	165.75	25
	$s=2$	.55	290.51	84	$s=2$	.65	427.25	52
	$s=3$	.82	715.63	129	$s=3$	.92	1066.8	81
	$s=4$	.93	1365.9	176				
	$s=5$	.95	2125.5	225				
		$\tilde{\lambda}_i$	$-T\sum\ln(1-\tilde{\lambda}_i)$	df		$\tilde{\lambda}_i$	$-T\sum\ln(1-\tilde{\lambda}_i)$	df
83M3-97M4	$s=1$	.46	104.98	41	$s=1$	.46	106.78	25
	$s=2$	.58	254.15	84	$s=2$	.70	310.35	52
	$s=3$	.79	525.35	129	$s=3$	.91	724.53	81
	$s=4$	.93	985.98	176				
	$s=5$	.97	1601.27	225				
		$\tilde{\lambda}_i$	$-T\sum\ln(1-\tilde{\lambda}_i)$	df		$\tilde{\lambda}_i$	$-T\sum\ln(1-\tilde{\lambda}_i)$	df
86M9-97M4	$s=1$	.57	108.45	41	$s=1$	.43	73.28	25
	$s=2$	.65	241.95	84	$s=2$	.58	183.81	52
	$s=3$	.84	477.95	129	$s=3$	.93	527.30	81
	$s=4$	.94	836.84	176				
	$s=5$	.96	1267.1	225				

Table 4: Common Feature Tests Statistics

		Cod(1)	*Cod(2)*	*Cod(3)*	*Cod(4)*	*Cod(5)*	*Cod(6)*	*Cod(7)*	*Cod(8)*
OCA 5	$s = 1$	.005	.086	.628	.605	.509	.783	.911	.933
	$s = 2$	<.001	<.001	.001	.191	.330	.186	.266	.740
	$s = 3$	<.001	<.001	<.001	.002	.007	.003	<.001	<.001
	$s = 4$	<.001	<.001	<.001	<.001	<.001	<.001	<.001	<.001
	$s = 5$	<.001	<.001	<.001	<.001	<.001	<.001	<.001	<.001
OCA 3	$s = 1$	<.001	.023	.334	.336	.301			
	$s = 2$	<.001	<.001	.009	.032	.114			
	$s = 3$	<.001	<.001	<.001	<.001	<.001			

Table 5: Testing for Non-Synchonous cycles

codependence vector after a adjustment of b months.

For the group of five countries, only two codependence vectors are found. For the smaller group with Germany, Belgium and The Netherlands, two codependence vectors and thus one common cycle emerges with a delay of adjustment of five months. These two vectors are $Nl - .478Ger$ and $Be - 1.811Ger$. However the determination of this reaction time remains arbitrary. Whether it is sustainable to form a union if idiosyncratic shocks need one year or five months to be fully absorbed remains a matter of judgment. Also a few number of common feature vectors is found when we enlarge the number of countries. By contrast, our analysis based on the Weak Form Reduced Rank Structure showed that short-run fluctuations are not perfectly synchronized.

Compared to the rest of the empirical literature, these results yield two pieces of information. Firstly, they confirm the idea of a two speed Europe, and suggest that even a restricted monetary union would still face some stabilization cost. Of course, whether this cost is outweighed or not by the gains obtained from adopting a unique currency remains beyond the scope of this paper. Secondly, we show that allowing for a multivariate empirical concept of an OCA does challenge some findings of the literature, i.e. that some hard core of the EMU including Germany and its small neighbors is a robust feature.

Furthermore, our results cocerning the non synchronous common features allow to shed an interesting light on the findings of Rubin and Thygesen (1996). It is worth reminding that their results support the idea of a costless monetary union made of at least nine countries. The codependence order (the b measure) obtained by these authors point out an adjustment time of 2 or 3 months. Our results show that their findings can be explained in two ways. (i) As shown by Hecq (1998) and Cubbada (1999), the use of seasonally adjusted variables may lead to the detection of spurious common features vectors and so to an underestimation of the number of common cycles. Our analysis of non synchroneous common cycles with non adjusted data lead to a similar result but with an adjustment time quite higher. (ii) The striking difference on the diagnostics in terms of OCA stems from the different econometric methodiologies; if one allows for adjustment delays like in Rubin in Thygesen (1996) Europe and probably EMU constitute an OCA; if one does not, no OCA is found; in this paper, we have argued that the choice of a maximum adjustment time is arbitrary and does not allow to determine the OCA without some *a priori* judgement.

From a quite paradoxical point of view, our results tend to be more useful for the analysis of the current situation (i.e. 11-countries EMU) than for the answer to the outdated debate on the optimality of the European Monetary Union. First, as said before, our concept of perfect OCA leads to a very strict criterion. From the cost side of the monetary union, some degree of potential stabilization -although quite limited in our view- is still possible. Furthermore, our analysis ignores the benefits of EMU that have to be balanced against the costs in order to define an OCA. Nevertheless, these results (but also the one related to the non synchronous approach) may shed an interesting insight on the potential problems encountered to conduct a single monetary policy in Europe. The recent experience of the European Central Bank shows that some desynchronisation in the national cycles between EMU members can make the choice of the timing and the stance of monetary policy quite difficult[19]. The recent hike in the short-run interest rates undertaken by the ECB[20] has been assessed quite differently across countries : for some EMU members like the Netherlands or Spain[21] which faced a booming economic activity and inflationary pressures, it came too late; for other members like Germany and Italy, the tightening may hamper their very recent recovery from a (hard) recession and comes to a certain extent too early; finally, for countries like Belgium or France, the timing appeared optimal. In this respect, our results emphasize the non existence of a unique common cycle, even between countries thought to be highly synchronized and are thus fully consistent with the recently observed divergences. The same holds for the results of the non synchronous approach

[19]This is different from another major problem related to the different transmission mechanisms of the monetary policy. This last problem points out the difference in the structures of the economy while our argument refers to the business cycles.

[20]On November 4, the ECB raised the rate on the main refinancing operation in the Eurosystem from 2.5% to 3%.

[21]In late April, in the absence of a tight monetary policy, the Spanish government cut some prices in the electricity and telecommunications sectors to cool off the economy.

that emphasizes an adjustment delay of five months.

5 Conclusion

The aim of this paper is to assess the composition of the European OCA on the basis of the degree of asymmetry of real shocks. Identifying some weakness in the existing literature, we focus on a multivariate determination of the OCA and make use of the recent developments of the common feature analysis. We define a "Perfect OCA" as a set of countries for which the short-run fluctuations are perfectly synchronized. While maybe restrictive, this definition yields the advantage of clarity, unlike those allowing for "some sustainable reaction time".

Our results do not allow to isolate any Perfect OCA. They suggest that even a restricted monetary union composed by Germany, Belgium and The Netherlands could face a stabilization cost. The findings have some implications for the economic policy of the current Euro zone. In particular, they confirm that the optimal timing and the stance of the unique monetary policy will be difficult to define, given the adjutment delay between EMU countries. On the methodological side, our analysis points out the potential sources of divergent results with the other study resting on a multi country determination of an OCA, i.e. the one of Rubin and Thygesen (1996). In particular, it is suggested that using seasonally adjusted data may lead to find out spurious codependence vectors and that the non synchroneous approach yields quite different results from the synchroneous one. In this paper, we have argued that the synchroneous approach, while restrictive in some sense, has the advantage of non requiring any a priori idea on the "sustainable reaction time".

In order to assess the robustness of these results, one should carry out the same analysis on other proxy variables for economic activity. Nevertheless, given the fact that it is preferable to use non seasonally adjusted data and that the series have to be harmonized across countries, the number of potential candidates appears to be rather limited.

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